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Does Trade Increase Total Factor Productivity: Cointegration Evidence for Chile

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Abstract

In this study, we examine the long-run impact of capital goods imports, exports of manufactured and primary goods on total factor productivity in Chile. Using the integration and cointegration techniques of Kapetanios (2005), Pesaran, Shin, and Smith (2001), Stock (1987), and Saikkonen (1991) we find a long-run relationship between these variables. All in all, our estimation results provide evidence for the existence of productivity-enhancing effects of capital goods imports and manufactured exports and of productivity-limiting effects of primary exports.

Does international trade increase the productivity of an economy? This question has been the subject of considerable theoretical and empirical research in recent years. Theoretically, there are several arguments in favor of a long-run relationship between trade and aggregate productivity. With regard to imports, it can be argued that, in particular, capital goods imports may increase domestic productivity, since capital goods embody technological knowledge, and, hence, foreign R&D activities spill over from one country to another through international trade. The literature usually suggests two basic mechanisms for these trade-related R&D spillover effects (see, for example, Grossman and Helpman 1991, and Rivera-Batiz and Romer 1991). First, knowledge spillovers may arise from studying blueprints of new capital goods, so that the recipient countries can imitate the technology. If the cost of obtaining knowledge is less than the cost of the corresponding invention, then a knowledge spillover occurs. Second, if R&D expenditures by countries from abroad create new capital goods that are different or better than those that already exist, then the productivity of an importing country improves by employing a wider variety of capital inputs or by simply using better capital inputs in final production. Thus, capital goods imports enable a country to acquire technology developed worldwide.

Referring to exports, it can be argued that, first, an expansion in exports may promote specialization in sectors in which a country has comparative advantage, and lead to a reallocation of resources from the relatively inefficient non-trade sector to the more productive export sector. Second, the growth of exports can increase productivity by offering larger economies of scale (Helpman and Krugman 1985). And third, increasing exports may affect aggregate productivity through dynamic spillover effects on the rest of the economy (Feder 1983). The possible sources of these knowledge externalities include productivity enhancements resulting from increased competitiveness, more efficient management styles, better forms of organization, labor training, and knowledge about technology and international markets. As Chuang (1998) argues, entering competitive international markets requires knowledge about foreign buyer's specifications, quality and delivery conditions. To satisfy these requirements, foreign purchasers help and teach local exporters to establish each stage of the production process and improve management and marketing practices. The development of efficient quality control procedures, management and marketing methods, product specifications and production guidelines is simultaneously fostered by the competitive pressure in the international markets. In short, knowledge is generated through a systematic learning process due to exporting and spills over to the domestic economy.

In this context, the literature usually distinguishes between manufactured and primary exports. Lucas (1993), among others, points out that the dynamic technological spillover effects are mainly associated with manufactured exports rather than with primary exports. Furthermore, several authors hypothesize that primary exports are an obstacle to greater productivity growth. The main arguments advanced in support of this hypothesis are: (i) Primary products offer no sustainable potential for knowledge spillovers, and an increase in primary exports can draw resources away from the externality-generating manufacturing sector (Matsuyama 1992, Sachs and Warner 1995). (ii) Primary exports are subject to extreme price and volume fluctuations. Increasing primary exports may therefore lead to increasing GDP variability and macroeconomic uncertainty. High instability and uncertainty may, in turn, hamper efforts at economic planning and reduce the quantity as well as the efficiency of investments (Dawe 1996). Consequently, the theoretical literature suggests that the effects of exports on economic productivity differ significantly between primary and manufactured products.

Admittedly, the empirical literature on trade and productivity generally uses aggregate measures of international trade, such as, for example, the ratio of imports plus exports to GDP, usually referred to as *openness* (see, for example, Frankel and Romer 1999, Miller and Upadhyay 2000, Jonsson and Subramanian 2001, Alcalá and Ciccone 2004). Without exception, all studies appear to find a positive and statistically significant relationship between trade and total factor or labor productivity, where most of them are based on cross-country data. Few studies apply panel-data regressions (see, for example, Miller and Upadhyay 2000). To our knowledge, only Jonsson and Subramanian (2001) have investigated the relationship between trade and aggregate productivity using time series techniques.

What contribution can a further paper make to this already existing empirical literature? Many authors have addressed the methodological problems associated with country regressions, including simultaneity bias, parameter heterogeneity, and error autocorrelation; see the reviews by Levine and Renelt (1991) and the critique in Ericcsson, Irons, and Tryon (2001). Therefore, we apply time series techniques. More concretely, this paper applies a unit root test that allows for an unknown number of structural breaks. Moreover, we use cointegration analysis to examine the long-run impact of trade on total factor productivity. Thereby we check the robustness of our results by using two different methods to estimate the parameters of our long-run relation. In order to tackle the well-known simultaneity problem between trade and productivity, one approach taken in this paper ensures that our estimates are unbiased even if the explanatory variables are endogenous. As far as the measure of trade is concerned, we argue that aggregate trade measures may mask important differences between different trade categories. Therefore, we do not use the ratio of imports plus exports to GDP, but we disaggregate trade into the categories capital goods imports, manufactured exports, and primary exports.

In order to investigate the impact of capital goods imports, primary and manufactured exports on productivity, we use Chilean time series data from 1960 - 2001. Chile is an interesting case study for several reasons. First, Chile experienced a pattern of high long-run growth, which, however, was interrupted by the collapse of the Allende government in 1973, the 1975 recession, and the 1982 economic crisis. As shown in Chumacero and Fuentes (2005) total factor productivity played an important role in this growth process, especially in the period 1975 - 1981 and after 1985. Second, Chilean exports and imports grew very rapidly after 1974, when trade liberalization was initiated. Admittedly, this growth was interrupted by the balance-of-payments crisis in 1982. Third, Chilean exports rely heavily on primary products, although the share of manufacturing exports in goods exports rose from about 7 percent in 1973 to 44 percent in 2001. And fourth, as found, for example, by Romaguera and Contreras (1995), Chile is extremely vulnerable to fluctuating commodity prices, especially copper prices. Nowadays, copper still accounts for about 37 percent of total exports of goods in Chile. [2]

The rest of the paper is organized as follows. Section 1 presents the empirical model and the data. The econometric methodology is described in Section 2. The estimation results are presented in Section 3. A final section summarizes the conclusions.

1. Empirical model and data

1.1. Model

Or objective is to investigate if and how increasing capital goods imports, manufactured exports and primary exports affect total factor productivity. To keep the analysis simple, we assume that the level of total factor productivity, TFP_t, at time t can be expressed as a Cobb-Douglas function of capital goods imports, CM_t, manufactured exports, IX_t, primary exports PX_t, and other exogenous factors C_t:

$$TFP_t = C_t \, CM_t^{\alpha} \, IX_t^{\beta} \, PX^{\delta} \,, \tag{1}$$

where α, β, and δ are the elasticities of TFP_t with respect to CM_t, IX_t, and PX_t. Taking natural logs (ln) of both sides of equation (3) gives the estimable linear function

$$\ln TFP_t = c + \alpha \ln CM_t + \ln IX_t + \delta \ln PX_t + e_t, \tag{2}$$

in which all coefficients are elasticities, c is a constant parameter, and e_t is the usual error term, which reflects the influence of all other factors. Since equation (2) can be interpreted as the long-run (cointegration) relationship between the variables, we estimate α, β, and δ using cointegration techniques in order to determine the long-run impact of increasing capital goods imports, manufactured exports and primary exports on total factor productivity in Chile.

Before estimation, we follow Miller and Upadhyay (2000) and Jonsson and Subramanian (2001) and calculate lnTFP_t on the basis of a constant-returns-to-scale Cobb-Douglas production function,

$$Y_t = TFP_t K_t^{a} L_t^{(1-a)}, \tag{3}$$

where Y_t denotes the aggregate production of the economy, and TFP_t, K_t, L_t are the level of total factor productivity, the capital stock, and the stock of labor, respectively. The key parameter necessary for the calculation are the factor-output elasticities a and (1 - a). From the growth accounting point of view, these elasticities are given by the capital and labor shares from the national accounts. We calculated the average capital share in the period 1960 - 2001 to be a = 0.5, which is in line with the results of Chumacero and Fuentes (2005) who report an average value of a = 0.507 for the period 1960 - 2000. Therefore, we set a = 0.5 to calculate the natural logarithm of total factor productivity [1] according to

$$\ln TFP_t = \ln Y_t - 0.5 \ln K_t - 0.5 \ln L_t . \tag{4}$$

1.2. Data

The data used to calculate $\ln TFP_t$ and to estimate equation (2) are annual from 1960 to 2001. They were gathered from the *Indicadores económicos y sociales de Chile 1960-2000* and the *Boletínes mensuales* published by the Chilean Central Bank. The output, Y_t, is measured by real Chilean GDP. K_t is the Chilean capital stock in real terms, which was computed on the basis of accumulated capital expenditure using the perpetual inventory method. GDP, capital stock, capital goods imports, exports of manufactured products, and primary products are evaluated in Chilean pesos at constant 1996 prices. The labor variable, L_t, is represented by the total number of people employed each year. Figure 1 shows the evolution of total factor productivity, capital goods imports, exports of manufactured products, and primary products in the period under consideration. (All variables are in logarithms).

Plot of Time Series

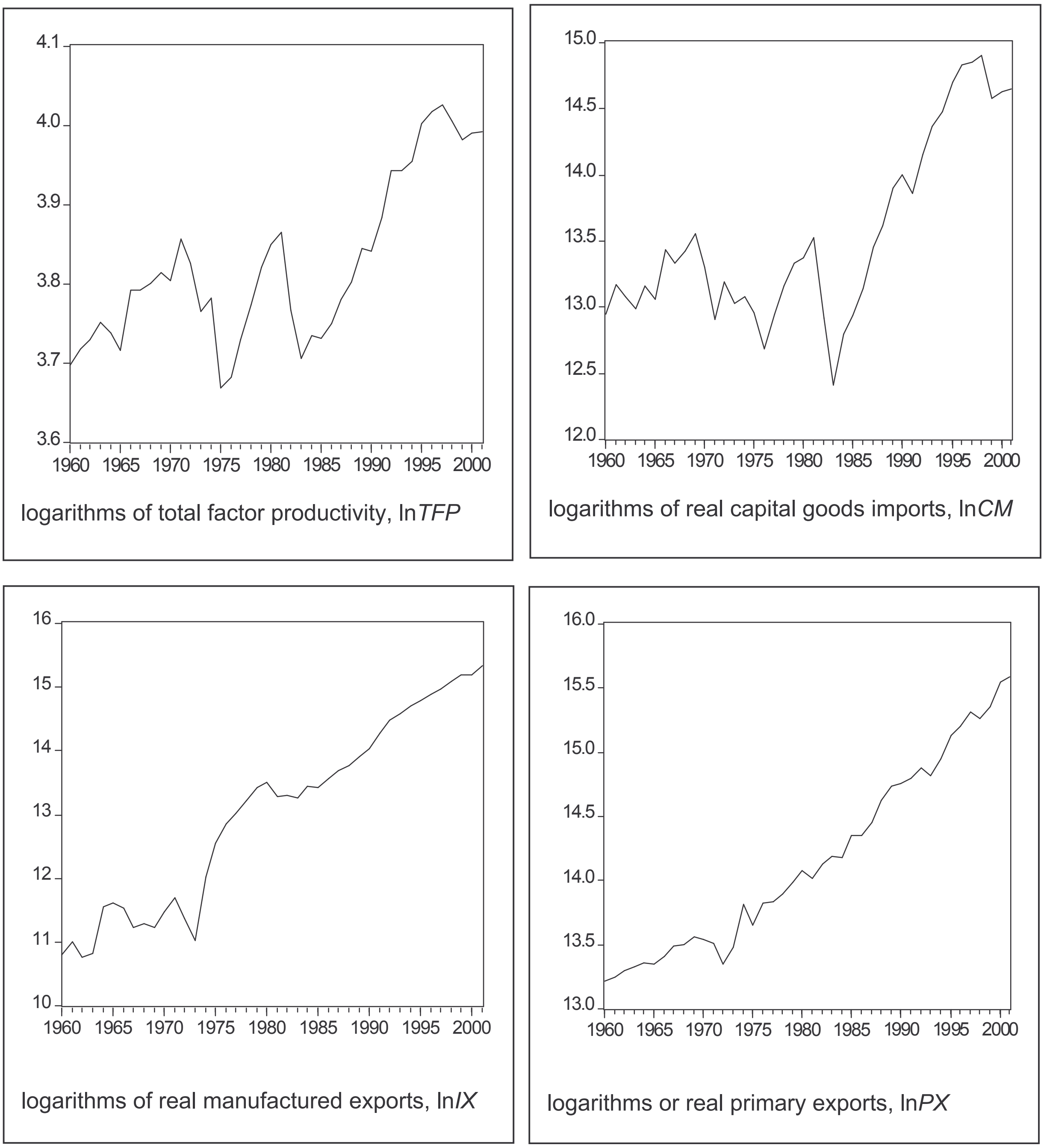

*Figure 1.*Time series used

2. Methodology

2.1. *Time series properties*

In the first step, we test the time series lnTFP_t, lnCM_t, lnIX_t, and, lnPX_t for unit roots. It is well known that standard unit root tests are biased in favor of identifying data as integrated if there are structural changes. For all the series there is indeed a strong likelihood that structural discontinuities are present (for example, the socialist government of President Allende (1970-1973), the drastic trade liberalization initiated in 1974 by the military Regime of Pinochet, the 1975 economic crisis, and the deep 1982 recession). Therefore, we use the unit root test recently developed by Kapetanios (2005). The Kapetanios procedure permits a formal evaluation of the time series properties in the presence of structural breaks at unknown points in time. It allows the break dates to be identified endogenously through the testing procedure itself. In order to test the unit-root hypothesis against the alternative of one or more structural breaks, we estimate three models of the Dickey-Fuller type without any prior knowledge of any potential break dates:

$$y_{1t} = \mu_1 + b_1 t + a_1 y_{t-1} + \sum_{i=1}^{m} \delta_1 DU_{i,t} + \sum_{i=1}^{k} c_{1i} \Delta y_{t-1} + e_{1t}\,, \tag{5}$$

$$y_{2t} = \mu_2 + b_2 t + a_2 y_{t-1} + \sum_{i=1}^{m} \delta_2 DT_{i,t} + \sum_{i=1}^{k} c_{2i} \Delta y_{t-1} + e_{2t}\,, \tag{6}$$

$$y_{3t} = \mu_3 + b_3 t + a_3 y_{t-1} + \sum_{i=1}^{m} \delta_3 DU_{i,t} + \sum_{i=1}^{m} \delta_3 DT_{i,t} + \sum_{i=1}^{k} c_{3i} \Delta y_{t-1} + e_{3t}\,, \tag{7}$$

where y_{it}, i = 1, 2, 3, is the variable considered (lnTFP_t, lnCM_t, lnIX_t, lnPX_t), k is the lag length, t is the time trend, m denotes the number of structural breaks, and $DU_t = 1$ (t >TB), $DT_t = 1$ (t > TB)(t -TB) are indicator dummy variables for the break at time TB $\in$ T (T = 42, 1 ≤ t ≤ 42).

As can be seen from the equations, Model (5) allows for up to m changes in the intercept of the trend function, while Model (6) accounts for m breaks in the slope of the trend function without a change in the level. Model (7) allows for both effects to take place simultaneously.

In the empirical analysis we consider the possibility that up to two break points occurred over the relevant period: m = 1, m = 2. The first break point is chosen by estimating the models for each possible break date in the data set, and TB is selected as the value which is associated with the minimum sum of squared residuals ($\sum(e_{1t})^2, \sum(e_{2t})^2, \sum(e_{2t})^2$). If the corresponding minimum t-statistic of the hypothesis $a_i = 1$ under model i = 1, 2, 3 does not exceed (in absolute value) the critical value reported by Kapetanios (2005), the unit root hypothesis is not rejected. Imposing the estimated break date on the sample, we start looking for the second break. Again, the second break point is chosen with the minimum sum of squared residuals in order to test the null hypothesis $a_i = 1$.

2.2. Cointegration

The second step is to test for cointegration. We apply the autoregressive distributed lag (ARDL) approach developed by Pesaran, Shin, and Smith (2001), which is also known as the bounds test procedure. The ARDL method has at least two advantages over alternatives such as the Engle and Granger (1987) two-step estimation procedure or the maximum likelihood approach developed by Johansen (1995). First, it can be applied to studies that have a small sample size, whereas the Engle-Granger and the Johansen procedures are not reliable for small sample sizes, such as in the present study. Second, the bounds test approach is applicable irrespective of whether the explanatory variables are I(1) or I(0). The test procedure is based on an unrestricted error correction model, which in our case is given by

$$\Delta \ln TFP_t = \alpha_0 + \delta_1 \ln TFP_{t-1} + \delta_2 \ln CM_{t-1} + \delta_3 \ln IX_{t-1} + \delta_4 \ln PX_{t-1}$$
$$\sum_{i=1}^{k} \beta_i \Delta \ln TFP_{t-i} + \sum_{i=0}^{k} \gamma_i \Delta \ln CM_{t-i} + \sum_{i=0}^{k} \lambda_i \Delta \ln IX_{t-i} + \sum_{i=0}^{k} \theta_i \Delta \ln PX_{t-i} + \mu_t \quad . \qquad (8)$$

In this model, which can be interpreted as an autoregressive distributed lag model, we test the absence of a long-run relationship between lnTFP_t, lnCM_t, lnIX_t, and lnPX_t by calculating the F-statistic for the null of no cointegration $H_0 : \delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$ against the alternative $H_1 : \delta_1 \neq \delta_2 \neq \delta_3 \neq \delta_4 \neq 0$.

The distribution of the test statistic under the null depends on the order of integration of the variables. In the case where (a) all four variables are I(0) the asymptotic 1% critical value is 4.29 - see Pesaran, Shin, and Smith (2001, p. 300, Table CI(iii)). If the calculated F-statistic falls below this value, the null hypothesis cannot be rejected (at the 1% significance level). In the case where (b) one or more series are I(0) and one or more series are I(1), the critical value falls in the interval [4.56, 5.61]. If the F-statistic falls within this bounds, the result is inconclusive. Thus, the order of integration must be known before any conclusion can be drawn. In the case where (c) lnTFP_t, lnCM_t, lnIX_t, and lnPX_t are I(1), the 1% critical value is 5.61. If the F-statistic lies above 5.61, the null of no cointegration is rejected at the 1% level.

2.3. Long-run elasticities

If lnTFP_t, lnCM_t, lnIX_t, and lnPX_t are found to be cointegrated, we use the simple Stock (1987) approach, which involves estimating equation (8) in order to obtain α, β, and δ. Given the low frequency of the data (annual) and the small sample size, we set the lag length k equal to 2. Then, following Lütkepohl and Wolters (1998), the general-to-specific approach is used by removing insignificant variables step by step until there remain only coefficients significant at the 1% level. Normalizing on the total factor productivity variable, or by δ_1, gives the long run relationship between capital goods imports, manufacturing exports, primary exports and total factor productivity.

However, if the explanatory variables are not weakly exogenous, the resulting estimates may be biased and inefficient and t-tests based on the model parameters may be misleading. Therefore, we check the robustness of the long-run estimates by means of the Dynamic OLS (DOLS) procedure developed by Saikkonen (1991). This procedure is asymptotically equivalent to Johansen's maximum likelihood estimator and is known to perform well in small samples. Moreover, DOLS has been shown to provide unbiased and asymptotically efficient estimates, even in the presence of endogenous regressors (Stock and Watson 1993). The DOLS regression in our case is given by equation (9) below:

$$\begin{aligned} \ln TFP_t &= c + \alpha \ln CM_t + \beta \ln IX_t + \delta \ln PX_t \\ &\quad + \sum_{i=-k}^{i=k} \Phi_1 \Delta \ln CM_{t+i} + \sum_{i=-k}^{i=k} \Phi_2 \Delta \ln IX_{t+i} + \sum_{i=-k}^{i=k} \Phi_3 \Delta \ln PX_{t+i} + \varepsilon_t \end{aligned} , \qquad (9)$$

where α, β, and δ are the long-run elasticities we are interested in, and Φ_1, Φ_2, and Φ_3 are coefficients of lead and lag differences of the I(1) variables, which are treated as nuisance parameters. They serve to adjust for possible endogeneity, autocorrelation, and non-normal residuals and result in consistent estimates of α, β, and δ.

3. Empirical Results

3.1. Unit root test results

Table 1 reports the results of testing the unit root null against the alternative of *m* structural breaks. At first, it can be seen that the estimated break points coincide with the beginning of the government of Allende in 1970, the drastic trade liberalization initiated by the military government of Pinochet in 1974, the 1975 economic crisis, and the 1982 recession. Furthermore, the test statistics show that we generally cannot reject the unit root hypothesis in favor of broken trend stationary at the 1% level. Since for the first differences the unit root hypothesis can be rejected, we can conclude that $\ln TFP_t$, $\ln CM_t$, $\ln IX_t$, and $\ln PX_t$ are integrated of order one, I(1).

However, using model (7) with one or two structural breaks, the unit root null can be rejected at the 5% level for the log of manufactured and primary exports. Accordingly, $\ln IX_t$, and $\ln PX_t$ are possibly trend stationary with at least one structural break. This introduces a degree of uncertainty to the analysis. Therefore, in the next step, we use the bounds test approach to the problem of testing for the existence of a long-run relationship between capital goods imports, manufacturing exports, primary exports and total factor productivity. As already mentioned, this approach is applicable irrespective of whether the explanatory variables are I(1) or I(0).

Table 1. Kapetanios (2005) unit root test

		m = 1			*m* = 2		
Series	Model	Dummy variable (Break year)	Test statistic	Critical value 5% (1%)	Dummy variables (Break years)	Test statistic	Critical value 5% (1%)
Levels							
$\ln TFP_t$	(5)	*DU*75 (1974)	-3.39	-4.930 (-5.338)	*DU*75, *DU*82 (1974, 1981)	-4.41	-5.685 (-6.162)
$\ln CM_t$	(5)	*DU*71 (1970)	-2.75	-4.930 (-5.338)	*DU*71, *DU*82 (1970, 1981)	-3.70	-5.685 (-6.162)
$\ln IX_t$	(5)	*DU*74 (1973)	-3.94	-4.930 (-5.338)	*DU*74, *DU*82 (1973, 1981)	-3.95	-5.685 (-6.162)
$\ln PX_t$	(5)	*DU*71 (1970)	-3.87	-4.930 (-5.338)	*DU*71, *DU*81 (1970, 1980)	-4.01	-5.685 (-6.162)
$\ln TFP_t$	(6)	*DT*75 (1974)	-2.45	-4.495 (-5.014)	*DT*75, *DT*82 (1974, 1981)	-2.41	-5.096 (-5.616)
$\ln CM_t$	(6)	*DT*82 (1981)	-3.95	-4.495 (-5.014)	*DT*82, *DT*75 (1981, 1974)	-3.76	-5.096 (-5.616)
$\ln IX_t$	(6)	*DT*81 (1980)	-3.34	-4.495 (-5.014)	*DT*81, *DT*74 (1980, 1973)	-3.95	-5.096 (-5.616)
$\ln PX_t$	(6)	*DT*74 (1973)	-4.88	-4.495 (-5.014)	*DT*74, *DT*82 (1973, 1981)	-4.92	-5.096 (-5.616)
$\ln TFP_t$	(7)	*DU*75 *DT*75 (1974)	-3.47	-5.081 (-5.704)	*DU*75, *DU*82 *DT*75, *DT*82 (1974, 1981)	-5.89	-6.113 (-6.587)
$\ln CM_t$	(7)	*DU*82 *DT*82 (1981)	-4.06	-5.081 (-5.704)	*DU*82, *DU*75 *DT*82, *DT*75 (1981, 1974)	-4.58	-6.113 (-6.587)
$\ln IX_t$	(7)	*DU*74 *DT*74 (1973)	-5.24	-5.081 (-5.704)	*DU*74 *DU*71 *DT*74 *DT*71 (1973, 1970)	-6.20	-6.113 (-6.587)
$\ln PX_t$	(7)	*DU*71 *DT*71 (1970)	-5.32	-5.081 (-5.704)	*DU*71, *DU*81 *DT*71, *DT*81 (1970, 1980)	-6.43	-6.113 (-6.587)
First differences							
$\Delta(\ln TFP_t)$	(5)	*i*75	-6.95	-3.53 (-4.23)	*i*75 *D*82	-6.87	-3.53 (-4.23)
$\Delta(\ln CM_t)$	(5)	*i*71	-6.31	-3.53 (-4.23)	*i*71 *D*82	-6.83	-3.53 (-4.23)
$\Delta(\ln IX_t)$	(5)	*i*74	-5.19	-3.53 (-4.23)	*i*74 *D*82	-4.89	-3.53 (-4.23)
$\Delta(\ln PX_t)$	(5)	*i*71	-8.39	-3.53 (-4.23)	*i*71 *D*81	-8.37	-3.53 (-4.23)

Notes: The lag length was chosen using the Akaike information criterion. The dummy variables are specified as follows: *i*71, *i*74, *i*75, *i*81, *i*82 are impulse dummy variables with zeros everywhere except for a one in 1971, 1974, 1974, 1981, 1982. *DU*71, *DU*74, *DU*75, *DU*81, *DU*82 are 1 from 1971, 1974, 1974, 1981, 1982 onwards and 0 otherwise. *DT*71 (*DT*71, *DT*75, *DT*81, *DT*82) is 0 before 1971 (1974, 1975, 1981, 1982) and *t* otherwise. Critical values for the levels are provided by Kapetanios (2005). Critical values for the first differences are from MacKinnon (1991). For the first differences, only impulse dummy variables were included in the regression. Impulse dummy variables, that is, those with no long-run effect, do not affect the distribution of the MacKinnon Test statistics.

3.2 Cointegration test results

In order to conduct the cointegration test developed by Pesaran, Shin, and Smith (2001) we need to establish the lag order, k, of the unrestricted error correction model (8). The lag order is selected using the Akaike (AIC) information criterion, the Hannan-Quinn (HQ) criterion, and the Schwarz (SC) criterion with a maximum of three lags. Table 2 shows the results of the lag selection procedure. As can be seen, both the Akaike information and the Hannan-Quinn criterion chose a lag length of two, whereas the SC selects a lag order of one.

Table 2. Lag length selection

k	AIC	HQ	SC
0	-0.074	0.266	0.048
1	-6.162	**-5.139**	-5.795
2	**-6.836**	-5.130	**-6.224**
3	-6.704	-4.315	-5.847

Notes: k is the lag order of the underlying VAR model for the conditional error correction model. Bold indicates lag order selection by the criterion.

Therefore, we estimate the ARDL model with one and two lags. Furthermore, an impulse dummy for 1971, $i71$, is included in equation (8) to make the residuals normally distributed. One possible reason for $i71$ to be important are the drastic reforms of the Allende government in that year (the nationalization of the mining, banking and agricultural sectors along with a expansionary fiscal policy). The results of the F-tests based on the unrestricted error correction model with k = 1 and k = 2 lags are reported in Table 3. Since the calculated F-statistics exceed the upper bound of the critical value band, we reject the null hypothesis of no long-run relationship between capital goods imports, manufacturing exports, primary exports and total factor productivity at the 1% significance level.

Table 3. F-statistics for testing the existence of a long-run relationship

k	*F*-statistic	1% critical value bounds of the F-statistic	
		I(0)	I(1)
1	16.35	4.29	5.61
2	5.89	4.29	5.61

Notes: k denotes the lag length. The critical value bounds are from Table CI(iii) in Pesaran, Shin, and Smith (2001, p. 300) with 3 regressors.

This result has a further implication: It justifies considering $\ln IX_t$, and $\ln PX_t$ as I(1) variables, because the F-statistics are higher than the 1% critical value, if all series are I(1). Therefore, in the following we treat all variables as I(1) when estimating α, β, and δ.

3.3 Estimation of the long-run coefficients: error correction model results

We use error correction model (8) to estimate the coefficients of the long-run relation between capital goods imports, manufacturing exports, primary exports, and the level of total factor productivity. That is to say, we regress $\Delta\ln TFP_t$ on $\ln TFP_{t-1}$, $\ln CM_{t-1}$, $\ln IX_{t-1}$, and $\ln PX_{t-1}$, the differences of $\ln CM_{t-1}$, $\ln IX_{t-1}$, and $\ln PX_{t-1}$ up to lag order two, the lagged difference of $\ln TFP_{t-1}$ also up to lag order two, an intercept term, and the impulse dummy *i*71. The following equation results by applying Hendry's general-to-specific approach, where successively the least significant variables are eliminated until there remain only coefficients significant at the 1%-level (t-statistics are given in parenthesis beneath the estimated coefficients):

$$\Delta \ln TFP_t = -0.691 \ln TFP_{t-1} + 0.120 \ln CM_{t-1} + 0.038 \ln IX_{t-1} - 0.087 \ln PX_{t-1}$$
$$\quad (-8.676) \qquad (8.404) \qquad (4.854) \qquad (-4.399)$$
$$+ 0.134 \Delta \ln CM_t - 0.041 \Delta \ln CM_{t-1} - 0.051 \Delta \ln IX_{t-1} + 1.754 + 0.120 i71 + \hat{e}_t$$
$$(10.874) \qquad (-3.245) \qquad (-4.646) \qquad (8.629) \quad (6.786) \qquad (10)$$
$$\overline{R}^2 = 0.869 \quad SE = 0.016 \quad JB = 1.501\,(0.472)$$
$$Arch(1) = 0.461(0.501) \quad Arch(2) = 0.208(0.813) \quad Arch(4) = 0.497(0.738)$$
$$LM(1) = 0.001(0.973) \quad LM(2) = 0.068(0.934) \quad LM(4) = 1.969(0.128)$$

The numbers in parentheses behind the values of the diagnostic test statistics are the corresponding *p*-values. These test statistics suggest that the model is well specified: The assumption of normally distributed residuals cannot be rejected (*JB*) and the Lagrange multiplier (*LM*) tests for autocorrelation based on 1, 2 and 4 lags, respectively, do not indicate any problems concerning autocorrelated residuals. The model also passes the LM tests for autoregressive conditional heteroscedasticity (*ARCH*) of order p = 1, 2, 4. Moreover, in Figure 2 CUSUM and CUSUM of square-tests are presented, which overall support a stable relation for the period of interest. Accordingly, the model does a good job even in the Chilean ‘breakdown periods’ (for example, 1975, 1982). Thus, statistically valid inferences can be drawn from the estimated model:

At first, we find that the coefficient of the lagged dependent variable in equation (10), $\ln TFP_{t-1}$, is negatively signed and highly significant, confirming that a long run relation exists among the variables. Furthermore, the t-statistics of $\ln CM_{t-1}$, $\ln IX_{t-1}$, and $\ln PX_{t-1}$ indicate that none of the variables could be excluded from the log-run relation. Normalizing on the coefficient of $\ln TFP_{t-1}$ finally yields the equation

Plot of Stability Tests

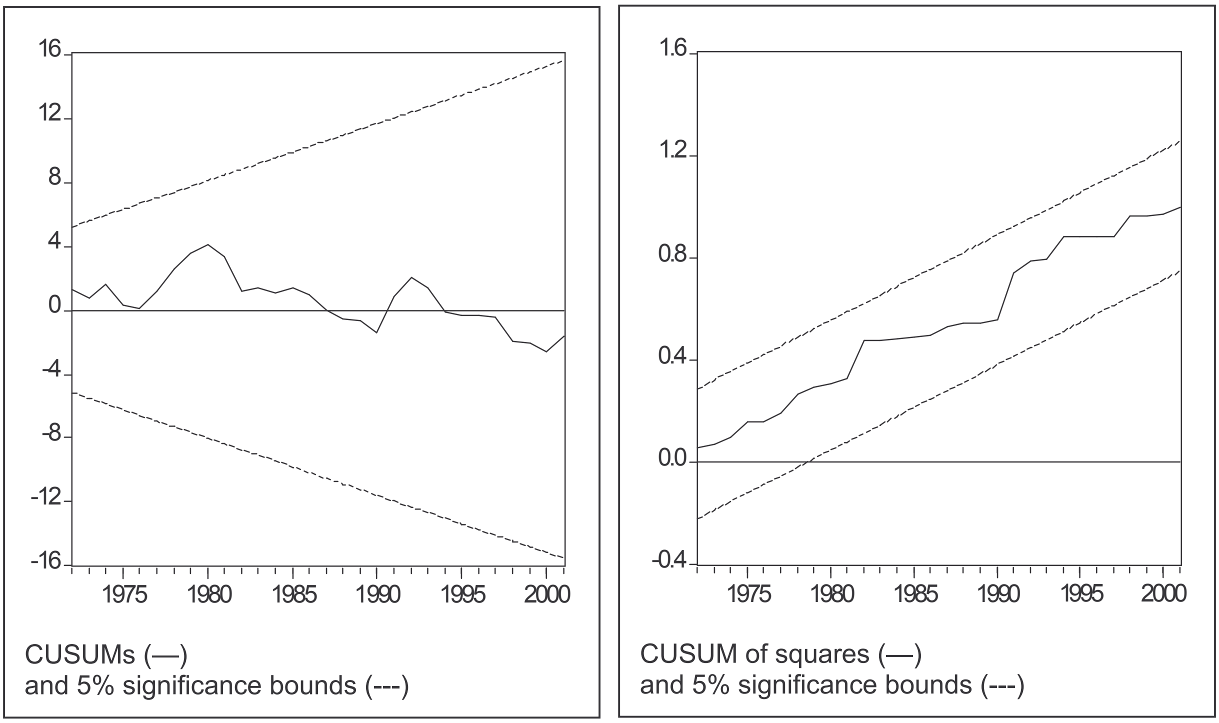

Figure 2. Stability tests

$$\ln TFP_t = 0.174 \ln CM_t + 0.056 \ln IX_t - 0.125 \ln PX_t \,. \qquad (11)$$

This equation implies that the level of total factor productivity increases by 0.174 percent in response to a one percent increase capital goods imports. Due to an increase of manufactured exports by one percent total factor productivity increases by 0.056 percent. In contrast, a one percent increase in primary exports leads to a 0.125 percent decrease of total factor productivity. Accordingly, capital goods imports are the main productivity determinant. However, due to endogeneity problems, regression (10) can by biased and the conclusions faulty. Therefore, in the final step of the analysis we check the robustness of the cointegration estimates.

3.4 Re-estimation of the long-run elasticities: dynamic OLS results

We re-estimate α, β, and δ by means of the DOLS procedure since DOLS generates unbiased and asympotically efficient estimates for variables that cointegrate, even with endogenous regressors. According to equation (9), we regress $\ln TFP_t$ on $\ln CM_t$, $\ln IX_t$, and $\ln PX_t$, the leads and lags of the differences of $\ln CM_t$, $\ln IX_t$, and $\ln PX_t$ up to order two, an intercept term, and the impulse dummy *i*71. The results of the DOLS procedure are presented in Table 4 (t-statistics are given in parenthesis beneath the estimated coefficients):

Table 4. DOLS procedure results

$\hat{\alpha}$	$\hat{\beta}$	$\hat{\delta}$
0.167 (12.203)	0.066 (3.894)	-0.132 (-3.163)

$\overline{R}^2 = 0.984 \quad SE = 0.018 \quad JB = 0.489\,(0.783)$

$Arch(1) = 0.657(0.423) \quad Arch(2) = 0.503(0.609) \quad Arch(4) = 0.554(0.649)$

$LM(1) = 0.003(0.954) \quad LM(2) = 0.795(0.468) \quad LM(4) = 2.282(0.112)$

The results are very similar to those of equation (10) and equation (11) respectively. Again, the diagnostic test statistics underneath Table 4 do not indicate any problems with autocorrelation, heteroscedasticity or nonnormality. All *p*-values exceed the usual (5%) significance levels. As in the error correction model estimation, the effects of capital goods imports and manufactured exports on total factor productivity are significantly positive (see $\hat{\alpha}$ and $\hat{\beta}$). The effect of an increase of primary exports on the level of total factor is again found to be strong and significantly negative (see $\hat{\delta}$). The magnitude of the coefficients in Table 4 does not differ substantially from equation (11). From this, we conclude that the coefficient estimates are fairly robust to different estimation techniques. Both the DOLS and the error correction model results can be interpreted as evidence of productivity-enhancing effects of capital goods imports and manufactured exports and of productivity-limiting effects of primary exports. This finding is in line with the theoretical consideration on the negative role that primary exports could exert on the productivity of an economy as discussed at the beginning of this paper.

4. Concluding remarks

This paper has used cointegration techniques to examine the impact of increasing capital goods imports, exports of primary products, and exports of manufactured goods on total factor productivity in Chile. The results suggest that there exists a long-run relationship between capital goods imports, manufactured exports, primary exports and total factor productivity. However, primary-product exports were found to have a statistically negative impact, whereas manufactured-product exports and capital goods imports have a statistically positive impact on total factor productivity. Accordingly, the estimation results are to be interpreted as evidence of productivity-enhancing effects of capital goods imports and manufactured exports and of productivity-limiting effects of primary exports. The latter may be due to the problem of fluctuating commodity export prices and earnings, especially copper prices, which is well known in the Chilean literature. Romaguera and Contreras (1995), for example, found that copper price volatility had negative effects on the Chilean economic development. Additionally, manufactured exports might offer greater potential for knowledge spillovers and other externalities than primary exports. Admittedly, the main productivity effects come from capital goods imports, suggesting that knowledge and technology is embodied in machinery and equipment and thereby transferred through international trade.

So, to return to the question posed at the beginning of this paper. Does trade increase the productivity of an economy? The Chilean evidence suggests that, all in all, increasing trade indeed raises productivity despite the productivity-limiting effects of primary exports. According to our results, the negative impact of primary exports is overcompensated by the positive effects generated by manufactured exports and capital goods imports. Furthermore, it is important to note that manufactured *and* primary exports provide the foreign exchange, that allows for increasing levels of capital goods imports to boost aggregate productivity.

Thus, the policy implications of our findings are ambiguous. One the one hand, it seems to be important to promote exports of manufacturing goods. On the other hand, it seems to be crucial to avoid trade-distorting measures that would counteract the comparative advantages that Chile has in primary exports, because such distortions can have the effect of reducing the capacity of financing capital goods imports.

Notes

1. Following the growth accounting literature, we also used the conventional value of the capital share of a = 0.333. The estimation results (not reported here, but available on request) do not change qualitatively when the total factor productivity is calculated based on a = 0.333.
2. In 1971 -1973 the share of copper represented almost 80% of total exports of goods, and the share of minerals as a whole announced to almost 90%. See, for example, Agosin (1999).

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